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Monitoring income-related health differences between regions in Great Britain: a new measurement framework

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Abstract

The paper proposes a new class of income-related health stratification indices that measure the extent to which differences in population health status between the regions of a country are systematically related to regional prosperity. The indices depend in general both on the degree to which the populations of different regions occupy well-defined layers or strata in the national distribution of the health outcome and on the scale of between-region differences in those outcomes if these are quantifiable, where the socioeconomic dimension is taken into account by ranking the regions in terms of economic prosperity rather than population health status. In particular, headcount and gap indices may be interpreted as measures of the overall incidence and depth of income-related health stratification between regions respectively, with the former well-defined for polytomous categorical variables without the need for etither dichotomisation or cardinalisation. The new measurement framework is used to examine the evolution of income-related health differences between the regions of Great Britain over the period from 1991 to 2008.

Keywords: income-related health stratification, regional analysis, ordinal data

JEL classifications: D63, I14, I18

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1. Introduction

Improvements in health over recent decades have not generally been matched by reductions in health inequalities (see e.g. van Doorslaer and Koolman 2004; Mackenbach et al., 2008; Marmot, 2013) leading to the identification of this issue as a key challenge in health care and social policy (Commission on Social Determinants of Health, 2008). As a result, trends in health inequalities are now regularly monitored in a number of countries utilising routinely collected administrative and survey data on population health outcomes reported at the local administrative unit or region level (Frank and Haw, 2011), with the specification of health inequality targets by policymakers at similar levels of spatial aggregation (see e.g. Department of Health 2001).

One limitation of this area-based approach to monitoring and policy formulation has been that it has focused almost exclusively on differences in overall population health between regions and thereby failed to take into consideration the variation in health outcomes between individuals within regions (Marmot et al., 2010). In particular, the social gradient for health outcomes that are cardinally measurable, such as mortality rates and life expectancy, may simply be interpreted as measures of between-region inequality based on population average health outcomes for each region. A second more general problem concerns the question of how to measure inequality using ordinal or categorical data, such as survey measures of self-reported health and subjective well-being, without first converting the data into cardinal form by assigning some more or less arbitrary numerical values either to each response category or to the differences between categories (see Allison and Forster, 2004; Lv et al, 2015; Kobus, 2015). The main objective of this paper is to develop an approach to measuring the extent to which differences in population health status between regions are systematically related to regional prosperity, which both takes account of intraregional variation in health outcomes and is directly applicable using either categorical or cardinal health data.

More specifically, we set out to gauge the degree of stratification between the population health distributions of the regions of a country, where this approach contrasts with the conventional focus in health inequalities research on "the evaluation of the inequality in the distribution of health status across individuals in a population" (Allison and Foster, 2004, p.505). The concept of stratification is deeply embedded within sociology, most notably in relation to the analysis of social class, but has only been of relatively recent concern within the economics literature. Thus Yitzhaki and Lerman (1991) in their seminal article quote a definition by the sociologist Lasswell (1965, p.10): "In its general meaning a stratum is a horizontal layer, usually thought of as between, above or below other such layers or strata. Stratification is the process of forming observable layers, or the state of being comprised of layers." Key to this definition is the idea that stratification, unlike segregation, implies a hierarchical ordering of groups according to some metric that, if cardinal, may be used to also measure the scale of the resultant differences in outcomes between groups. Accordingly, treating regional populations as groups, we seek to evaluate not only the degree to which the populations of different regions occupy well-defined layers or strata in the national health distribution but also the scale of between-region differences in health outcomes if quantifiable. The socioeconomic dimension is taken into account by ranking the regions in terms of economic prosperity rather than population health status.

Our approach builds on the class of univariate stratification indices introduced in Allanson (2015), which we employ directly to measure "pure" health stratification between regions. However, the main contribution of the paper is to provide an extension that leads to the specification of a class of income-related health stratification (IRHS) indices, which are related to the univariate indices in the same way that the health concentration index is to the health Gini. Thus we first rank regions by average income and then proceed to measure the income-related health stratification between each pair of regions as the product of an ‘identification index’ and an ‘alienation function’, where the terminology is borrowed from the analogous literature on polarisation (see Duclos et al., 2004). The identification index is defined as the difference in the probabilities that a randomly selected inhabitant of the richer region is more rather than less healthy than a randomly chosen inhabitant of the poorer region, where this measure captures the degree to which the two regions form well-defined strata in their combined health distribution. The alienation function is specified as a power function of the absolute difference in average health between the two regions, if quantifiable, with this being set equal to one by definition if the value of the power or exponent is set equal to zero. Finally the IRHS index is obtained by aggregating over all pairs of regions so as to yield a national population-weighted average of the pairwise indices.

As with Foster-Greer-Thorbecke (FGT) poverty measures (Foster et al., 1984), the choice of power in the alienation function determines the interpretation of the resultant IRHS indices. In particular, the zeroth power index is simply given as the population weighted mean difference in the probabilities that the health of a randomly chosen inhabitant of a richer region is better rather than worse than that of a randomly selected inhabitant of a poorer region, and is therefore easily understood as a measure of the income-related incidence of health stratification. A second major attraction of this ‘headcount’ index is that it requires no cardinal specification of categorical outcome data, providing a measure that is equal to twice the between-region generalised health concentration index for binary health status indicators but also directly applicable to polytomous categorical variables. Nevertheless, stratification is more than identification and, if suitable health data are available, it is also of interest to take alienation into account in order to obtain an index that fully captures the richness of the concept. The first power index, if defined, further takes into account absolute differences in average health between regions and may therefore be interpreted as a measure of the income-related health stratification 'gap' between regions. The IRHS gap index may be seen as an extension of conventional measures of inequality between regions, reducing to twice the between-region generalised concentration index if there is no overlapping of regional health distributions. Finally, second and higher integer power indices, although not considered further in this paper, will be directly sensitive to the distribution of absolute health differences between regions, being defined as convex functions of pairwise average health gaps.

Our methodology differs from most of the literature on the measurement of health inequality with categorical data in that it incorporates the socioeconomic dimension, with the seminal paper by Allison and Foster (2004) emphasising the point that their method is designed to evaluate overall inequality in health, without focusing on any particular cause or justification. One major exception is Zheng (2011) who proposes an approach similar to our own to the extent that it is based on a partion of the population into groups or classes that are ranked by income from poorest to richest. He proceeds to develop a set of welfare and inequality dominance conditions that could in principle be used to evaluate socioeconomic health inequalities between regions given this formal similarity between the two approaches. However Zheng (2011) narrows the definition of groups down to income quantiles, such that a member of a higher socioeconomic class will definitely be richer than one from a lower class, and the methodology would lose much of its normative significance if it was applied to groups, such as regional populations, with income distributions that exhibit a substantial degree of overlap. Makdissi and Yazbeck (2014) provide an alternative solution to the categorical data problem that exploits the availability in some surveys of information on multiple dimensions of health status in order to define a set of socioeconomic health inequality indices in the breadth (rather than the depth) of well functioning health attributes.

In contrast our methodology does not require multiple health indicators nor the censoring of the available information on (general) health status through the dichotomisation of polytomous variables.

The new class of indices is used to examine the evolution of income-related health differences between the regions of Great Britain. Current interest in the links between regional differences in health outcomes and deprivation within Britain may be traced back to the publication of the Black Report (Black et al., 1980). In particular, there has been a long-running debate (see Taulbut et al., 2013) about why health outcomes have been persistently worse in Scotland than in England and Wales even after controlling for differences in levels of social deprivation – the so-called 'Scottish' or 'Glasgow' effect. The ongoing impact of the financial crisis in 2008 has also renewed concerns about health differences between English regions, leading Public Health England to commission an independent inquiry on health equity for the North (Whitehead, 2014). This work has focused very largely on differences in life expectancy, with the evidence on regional disparities in self-reported measures of general health and disability both more limited and equivocal in nature. The current study sets out to identify income-related differences between regions in a range of self-assessed health measures available in the British Household Panel Survey (BHPS) over the period 1991 to 2008.

The paper is structured as follows. The next section introduces the proposed class of income-related health stratification indices and discusses the properties of the IRHS headcount and gap indices. Section 3 presents the empirical study of income-related health stratification in Great Britain. The final section summarises the contribution and offers some suggestions for further applications of the measures.

2. Definition and properties of the class of income-related health stratification indices

Consider the population of some country that is resident in $R \geq 2$ mutually exclusive and exhaustive administrative regions. The population and population share in region r $(r = 1,....R)$ are given as n_r and $p_r = n_r / N$ respectively, where $N = \sum_r n_r$ is the total size of the national population. Let H_r denote the health variable in region r with cumulative distribution function $F_r(h) = P(H_r \leq h)$ and inverse distribution or quantile function $F_r^{-1}(q)$.[2] The health distribution function for the national population is written as $F_u(h) = \sum_r p_r F_r(h)$ where $H_u = H_1 \cup H_2 \ldots \cup H_R$. The ranking of region r incomes in the region s and national health distributions are given as $F_s(F_r^{-1}(q))$ and $F_u(F_r^{-1}(q))$ respectively, with corresponding mean (fractional) ranks $\overline{F}_{rs}$ and $\overline{F}_{ru}$. We note that $\overline{F}_{rs} = \mathrm{P}(H_r > H_s)$ is the probability that the health of a randomly chosen resident of region r is better than that of a randomly chosen resident of region *s,* with $\overline{F}_{rr} = P(H_r > H_r) = 0.5$ by definition. Finally μ_r and $\theta_r = p_r \mu_r / \mu_u$, if defined, represent the mean health and health share of region r respectively, where $\mu_u = \sum_r p_r \mu_r$ is mean national health.

The measurement of stratification further requires the prior imposition of some ordering on the regions. In particular, in order to construct a univariate or 'pure' measure of health stratification regions first need to be ordered by population health status. Following Allanson (2015) this can be done by ordering regions by mean health with any ties separated on the basis of health distribution ranks such that $\mathrm{P}(H_s > H_r) > 0.5 > \mathrm{P}(H_r > H_s)$ for all

[2] $F_r(h)$ is assumed to be strictly increasing and continuous for notational convenience, implying that the probability of a randomly chosen resident of region r having the same health as a randomly selected resident of region s will have measure zero. The treatment of ties is discussed in the next sub-section.

relevant pairwise comparisons, where this secondary criterion could be used on its own if only ordinal health data were available such that mean health was not measurable.[3] However the main focus of this paper is on the construction of bivariate measures of income-related health stratification so in the remainder of this section regions are instead assumed to be ordered by mean income from the poorest $(r=1)$ to the richest region $(r=R)$, with ties separated as above on the basis of income distribution ranks when necessary. Income-related and 'pure' health stratification will be the same if income and population health status generate the same ordering. We note that small changes in individual incomes may lead to discontinuous changes in income-related health stratification if they lead to changes in the ordering of regions by income.

2.1 Measurement of pairwise income-related health stratification

The pairwise IRHS index $S_{rs}(\upsilon)$ between two regions *r* and *s* is taken to depend in general on both the degree to which the populations of the two regions occupy well-defined layers or strata in their combined health distribution and the scale of between-region differences in health outcomes. Specifically, we define $S_{rs}(\upsilon)$ as the product of an identification index I_{rs} and an alienation function $A_{rs}(\upsilon)$:

$$S_{rs}(\upsilon) = I_{rs}A_{rs}(\upsilon) \; ; \quad r,s = 1,....R \tag{1}$$

where the specification and interpretation of I_{rs} and $A_{rs}(\upsilon)$ are discussed in turn below.

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[3] The secondary criterion will generate a transitive ordering of a set of regions if the probability relationship between them exhibits mutual rank transitivity (see De Baets *et al.*, 2010), where the need for this condition arises iff there are more than two regions given that $\mathrm{P}(Y_s > Y_r) > 0.5$ and $\mathrm{P}(Y_t > Y_s) > 0.5$ does not necessarily imply $\mathrm{P}(Y_t > Y_r) > 0.5$. The empirical significance of the issue is likely to be limited but the condition can always be checked should the need arise.

The pairwise identification index I_{rs} in (1) is defined as:

$$
\begin{aligned}
I_{rs} &= \operatorname{sgn}(s-r)\big(\mathrm{P}(H_s > H_r) - \mathrm{P}(H_r > H_s)\big) \\
&= \operatorname{sgn}(s-r)\Big(\big(\mathrm{P}(H_s > H_r) + 0.5\,\mathrm{P}(H_s = H_r)\big) - \big(\mathrm{P}(H_r > H_s) + 0.5\,\mathrm{P}(H_s = H_r)\big)\Big) \\
&= 1 - 2\big(\mathrm{P}(H_r > H_s) + 0.5\,\mathrm{P}(H_s = H_r)\big)
\end{aligned}
\tag{2}
$$

where $$\operatorname{sgn}(s-r) = \begin{cases} 1 & \textit{if } s-r>0 \\ 0 & \textit{if } s-r=0 \\ -1 & s-r<0 \end{cases} \tag{2a}$$

I_{rs} is thus equal to the signed difference in the probabilities that a randomly chosen inhabitant of region s will have better rather than worse health than a randomly selected inhabitant of region *r*, where the use of the sign function ensures that $I_{rs} = I_{sr}$ and with $I_{rr} = I_{ss} = 0$ by construction. I_{rs} is defined for both continuous and discrete distributions, with the second line of (2) making explicit the treatment of ties in the case that $\mathrm{P}(H_r = H_s) \neq 0$, which will be the norm with self-reported health data from surveys in which individuals are asked to choose between a finite number of descriptive categories (e.g. very poor, poor, fair, good, excellent). In the empirical study, the convention is adopted that half of any ties are composed of pairs in which the inhabitant of the richer region is also healthier and half are pairs for which the opposite is the case, such that the net effect of any ties in the calculation of I_{rs} is zero. The final line of (2) follows by definition.

The interpretation of I_{rs} as an identification index follows from the observation that if individuals from the two regions are randomly matched with each other then I_{rs} will reflect the extent to which regional identity can be determined by assuming that the healthier individual within each pair will be from the richer rather than poorer region. I_{rs} will take its maximum value of one if regional identity can be determined with certainty by this rule, which will only be the case if there is complete separation of the populations of the two regions into discrete layers in their combined health distribution such that no inhabitant of the richer region is less healthy than any inhabitant of the poorer region: not only will everyone from the richer region be among the healthiest people in the two regions but also all the healthiest people will be from the richer region. Conversely, I_{rs} will equal zero if the health distributions of the two regions are identical such that the pairwise identification rule is entirely uninformative of regional identity: the healthier of any pair is equally likely to be from one region as the other if the two regions are indistinguishable in terms of health outcomes. Note that I_{rs} can also be negative, which will be the case if a randomly chosen inhabitant of the poorer region is likely to be healthier rather than unhealthier than a randomly chosen inhabitant of the richer region, taking a minimum value of minus one.

The alienation function in (1) is defined as:

$$A_{rs}(\upsilon) = \begin{cases} |\mu_s - \mu_r|^{\upsilon} & if\ \upsilon > 0 \\ 1 & if\ \upsilon = 0 \end{cases} \tag{3}$$

Alienation is thus specified as a power function of the absolute difference in population average health between the two regions if $\upsilon > 0$, and is set equal to one by definition if $\upsilon = 0$.[4] The parameter υ may be interpreted as an indicator of 'alienation' aversion. Thus a society in which the gap in health standards between the two regions is twice as large will have 2^{υ} times the level of alienation. Alternatively, υ is the elasticity of alienation with respect to the health standard gap, so that a 1% increase in the gap leads to a υ% increase in between-region alienation. For $\upsilon > 0$, $A_{rs}(\upsilon)$ will only be defined if it is possible to measure average health in which case it follows immediately that $A_{rs}(\upsilon) = A_{sr}(\upsilon)$ and $A_{rr}(\upsilon) = A_{ss}(\upsilon) = 0$. If $\upsilon = 0$ then the scale of the difference in population health status

[4] Following Allanson (2015) the specification of the alienation function could be elaborated by considering absolute differences in generalised or $\alpha-$order means between regions, with these interpretable as differences in 'representative health' levels between regions (cf. the definition of 'income standards' in Blackorby et al., 1981) with α being the Atkinson (1970) inequality aversion parameter. The special case considered here is obtained by setting $\alpha = 1$.

between regions is not of itself a matter of concern, with the index set equal to one irrespective of whether it is possible to actually measure mean health or not.

The parametric class of measures $S_{rs}(\upsilon) = I_{rs}A_{rs}(\upsilon)$ thus gives analysts and policymakers an instrument to evaluate regional health differences with varying sensitivity to distributional issues depending on social preferences. $S_{rs}(\upsilon)$ is symmetric in that $S_{rs}(\upsilon) = S_{sr}(\upsilon)$ but it is nevertheless sensitive to the ordering of regions by income, providing a 'directional' measure in the sense of Dagum (1997). In particular, in the limiting case of two regions with non-overlapping health distributions then $S_{rs}(\upsilon)$ will be positive if the health of the richer region *s* is better than that of the poorer region *r* and negative if the opposite is true.[5] Importantly, $S_{rs}(0)$ is well defined even if only ordinal health data are available, which is commonly the case when the source is a health or other general survey.

2.2. Definition and properties of the class of income-related health stratification indices

The class of income-related health stratification indices $S(\upsilon)$ is obtained as a population-weighted average of the pairwise indices $S_{rs}(\upsilon)$:

$$\begin{aligned} S(\upsilon) &= \sum_r \sum_s p_r p_s S_{rs}(\upsilon) = \sum_r \sum_s p_r p_s I_{rs} A_{rs}(\upsilon) \\ &= \begin{cases} \sum_r \sum_s p_r p_s \operatorname{sgn}(s-r)\left(\mathrm{P}(H_s > H_r) - \mathrm{P}(H_r > H_s)\right)\left|\mu_s - \mu_r\right|^{\upsilon} & \text{if } \upsilon > 0 \\ \sum_r \sum_s p_r p_s \operatorname{sgn}(s-r)\left(\mathrm{P}(H_s > H_r) - \mathrm{P}(H_r > H_s)\right) & \text{if } \upsilon = 0 \end{cases}; \end{aligned} \tag{4}$$

where $p_r p_s$ may be interpreted as the probability that the first of two individuals randomly selected with replacement from the national population will be from region *r* and the second from region *s*, and which therefore sum to one over all possible combinations.

$S(\upsilon)$ will take a value of zero if all pairwise indices $S_{rs}(\upsilon)$ are zero, although this does not necessarily imply that all regional health distributions are identical. $S(\upsilon)$ is strictly

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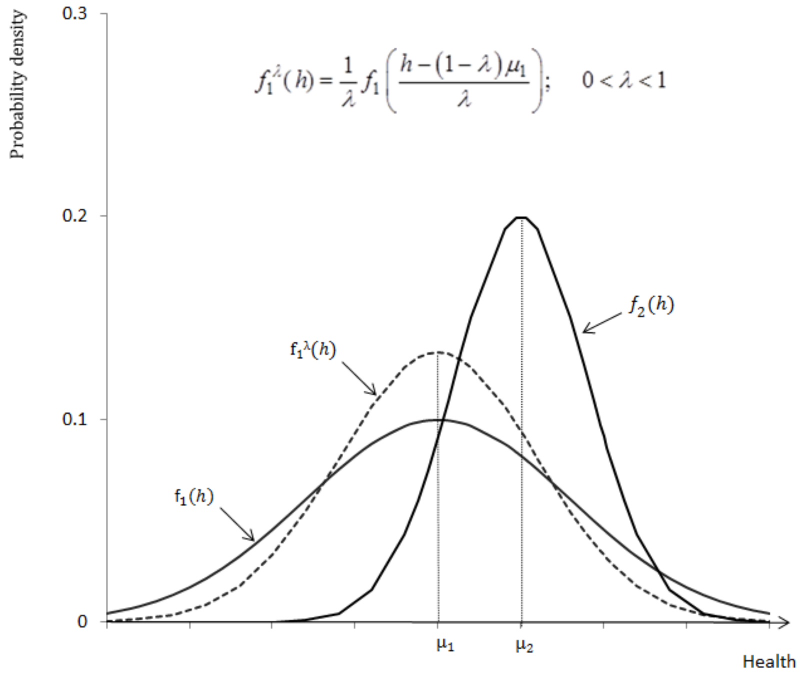
[5] The first case implies $(\mu_s - \mu_r) > 0$ and $I_{rs} = 1$ while the second implies $(\mu_s - \mu_r) < 0$ and $I_{rs} = -1$. As identification tends to zero in the two cases then $S_{rs}(\upsilon)$ will tend to zero from above and below respectively, changing sign when the identification index changes sign.

increasing in $S_{rs}(\upsilon)$, which provide unique estimates of the contribution of each pair of regions to income-related health stratification. Moreover, the pairwise indices may be meaningfully aggregated, given symmetry, to yield estimates $S_r(\upsilon) = p_r \sum_s p_s S_{rs}(\upsilon)$ of the contribution of each region to $S(\upsilon)$, with the further potential to identify the characteristics or factors that contribute to stratification. $S(\upsilon)$ is invariant to the permutation of regions and to the replication both of the subpopulations within regions (holding the population shares of the regions constant) and of the regions (holding the subpopulations within each region constant).

The dominance properties of $S(\upsilon)$ may be characterised in terms of identification and alienation axioms as in Allanson (2105), which provides a fuller discussion of the analogous properties of univariate stratification indices. First, if the country consists of two or more regions having symmetric, unimodal health densities with compact supports $f_r(h)$ and corresponding mean health levels μ_r then a symmetric, mean health preserving "squeeze" in the health distribution in any one region, say from $f_1(h)$ to $f_1^{\lambda}(h)$ as shown in Figure 1, cannot reduce identification and hence stratification if richer regions invariably have healthier populations (i.e. $\mu_s > \mu_r$ for all $s > r$ if mean health is defined). This property distinguishes stratification from inequality, since a reduction in within-region health variation holding between-region differences constant will lead to a fall in income-related health inequality according to the principle of health transfers (Bleichrodt and van Doorslaer, 2006). [6] Second, an identification-preserving scalar expansion of all health differences about national mean health μ_u, as illustrated by the national mean-preserving spread of population health levels in Figure 2, will unambiguously increase alienation and, if richer regions are also always healthier, stratification. Third, $S(\upsilon)$ will take its maximum value when the national

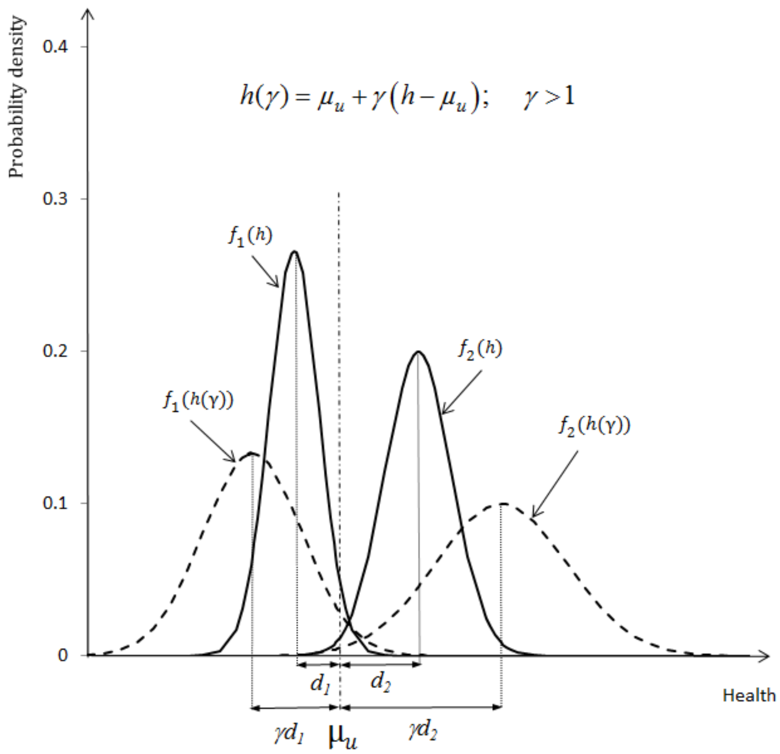
[6] This requires that inequality decreases if health is redistributed from a healthy individual to a less healthy one, leaving their respective ranking unchanged.

Figure 1. Illustration of identification dominance property



Note: $f_1^{\lambda}(h)$ has the same mean as $f_1(h)$ but is second-order stochastically dominant (see Duclos et al., 2004)

Figure 2. Illustration of alienation dominance property



population is equally divided between the two regions with the largest pairwise index $S_{rs}(\upsilon)$, where these may be expected to be the poorest and richest regions assuming a monotonic relationship between income and health outcomes. Finally, even though these properties mirror those of the social polarisation indices of Duclos et al. (2004) it is important to recognise that stratification is not the same as polarisation due to the different characterisations of identification employed in the two classes of measures.

2.3 Properties of the headcount index $S(0)$

The zeroth power index $S(0)$ may be re-written from (4) as:

$$\begin{aligned} S(0) &= \sum_r \sum_s p_r p_s I_{rs} = \sum_r \sum_s p_r p_s \operatorname{sgn}(s-r)\left(\mathrm{P}(H_s > H_r) - \mathrm{P}(H_r > H_s)\right) \\ &= 2\sum_r \sum_{s>r} p_r p_s \left(\mathrm{P}(H_s > H_r) - \mathrm{P}(H_r > H_s)\right); \end{aligned} \tag{5}$$

where $S(0)$ is the population-weighted average level of pairwise identification since $\sum_r \sum_s p_r p_s = 1$ by definition. More explicitly, $S(0)$ measures the average difference in the probabilities that the healthier of two randomly chosen individuals will come from the richer rather than the poorer region of which they are inhabitants.

Thus $S(0)$ may be interpreted as a headcount or incidence measure of income-related health stratification that captures the extent to which individuals' positions within the national health distribution are determined by regional prosperity. In particular, $S(0)$ will take its maximum value of $\left(1-\sum_r p_r^2\right)$ when residence in a particular region restricts individuals to a single interval or range of ranks in the health distribution that is exclusively occupied by inhabitants of their own region, with the ordering of the regions by population health status being the same as that by income. Dividing $S(0)$ by $\left(1-\sum_r p_r^2\right)$ yields a normalised index $\widehat{S}(0)$ that is the population-weighted average level of pairwise identification between all distinct regions with a maximum value of one. Conversely $S(0)=0$ if regional prosperity is entirely uninformative as a predictor of relative rank such that $I_{rs}=0$ for all pairs of regions, though a zero value may also arise in cases in which positive and negative values of the pairwise indices cancel each other out. Negative values of $S(0)$ imply that mean incomes by region are negatively correlated with population health outcomes.

$S(0)$ is a unit free measure that is continuous in individual health outcomes and invariant to rank-preserving transformations of them. If the health outcome measure is given by a binary indicator variable, taking values of zero and one, then $S(0)$ can be shown to be equal to twice the conventional between-region generalised concentration index since:

$$\begin{aligned} S(0) &= \sum_r \sum_s p_r p_s \operatorname{sgn}(s-r)\left(\mathrm{P}\left(H_s > H_r\right) - \mathrm{P}\left(H_r > H_s\right)\right) \\ &= \sum\nolimits_r \sum\nolimits_s p_r p_s \operatorname{sgn}(s-r)\left(P(H_s=1)(1-P(H_r=1)) - P(H_r=1)(1-P(H_s=1))\right) \\ &= \sum\nolimits_r \sum\nolimits_s p_r p_s \operatorname{sgn}(s-r)\left(P(H_s=1) - P(H_r=1)\right) \\ &= \sum\nolimits_r \sum\nolimits_s p_r p_s \operatorname{sgn}(s-r)\left(\mu_s - \mu_r\right) \\ &= 2\mu_u C_B \end{aligned} \tag{6}$$

where $\mu_u = \sum_r p_r \mu_r = \sum_r p_r P(H_r = 1)$ by definition and C_B is the between-region health concentration index. But, unlike the bewteen-region (generalised) concentration index, $S(0)$ is also defined for polytomous categorical variables without the need to first impose some essentially arbitrary cardinalisation of the health measure. Moreover, even in the dichotomous case $S(0)$ has a more natural and intuitive interpretation (cf. Wagstaff (2005) and the subsequent exchange of views following Erreygers (2009)).

With only two regions, the reduction in headcount IRHS $S(0)$ caused by a unit increase in one person's health would be greatest for inhabitants of the poorer region with health equal to the modal health level in the richer region, assuming that the poorer region also has worse population health. With more than two regions, the issue is more complicated as there is a need to consider which region to target as well as to identify which inhabitants of the targeted region to treat, where this will depend for intermediate regions on the net change in identification due to a unit change in the chosen person's health outcome. Nevertheless it is readily apparent that increasing the health of the poorest region, let alone the health of the least healthy inhabitants of that region, will not necessarily have the most impact on headcount IRHS: indeed $S(0)$ is invariant to changes in the health of inhabitants of the poorest region whose health is worse, and remains worse, than the most unhealthy inhabitant of any other region.

2.4 Properties of the IRHS gap index $S(1)$

The first power index $S(1)$ may be re-written from (4) as:

$$\begin{aligned} S(1) &= \sum_k \sum_l p_r p_s I_{rs} A_{rs}(1) = \sum_k \sum_l p_r p_s \left(\mathrm{P}(H_s > H_r) - \mathrm{P}(H_r > H_s)\right)(\mu_s - \mu_r) \\ &= S(0)\, A(1) + \operatorname{cov}\left(I_{rs}, |\mu_s - \mu_r|\right) \end{aligned} \tag{6}$$

where $A(1) = \sum_r \sum_s p_r p_s |\mu_s - \mu_r|$ is the average level of first-order alienation, i.e. the population-weighted mean of the absolute health gaps between regions; and $\operatorname{cov}\left(I_{rs}, |\mu_s - \mu_r|\right) = \sum_r \sum_s p_r p_s \left(I_{rs} - S(0)\right)\left(|\mu_s - \mu_r| - A(1)\right)$ may be interpreted as the covariance between pairwise levels of identification and alienation, which typically will be positive if health and income are positively correlated with each other at a regional level.

$S(1)$ is again interpretable as a population weighted average but the contribution that any particular pair of regions makes to the value of the overall index now depends not only on the pairwise identification index I_{rs} but also on the difference in mean health between them. For example, the lack of overlap between a healthy and an unhealthy region will count more towards the IRHS 'gap' as measured by $S(1)$ than the same lack between two regions both with moderate average health levels: in the limit, two regions with identical levels of mean health will not figure at all however large the difference in probabilities that a randomly chosen resident of one region will be better off than a randomly selected resident of the other.

$S(1)$ may also be seen to provide an extension of the conventional approach to the measurement of between-region inequality so as to additionally take account of the degree to which the populations of different regions occupy well-defined strata in the national health distribution. Thus $S(1)=2\mu_u C_B=\sum_r \sum_s p_r p_s \operatorname{sgn}(s-r)(\mu_s-\mu_r)$ if there is no overlapping of regional health distributions, i.e. the index reduces to twice the between-region generalised concentration index if there is full identification of all regional health distributions. Moreover $S(1)=2\mu_u G_B=A(1)=\sum_r \sum_s p_r p_s |\mu_s-\mu_r|$ in this case if income and health generate the same ordering of regions, where G_B is the between-region health Gini coefficient.

Dividing $S(1)$ by $A(1)$ yields the normalised IRHS gap index:

$$\hat{S}(1)=\sum_r \sum_s \frac{p_r p_s I_{rs}|\mu_r-\mu_s|}{A(1)}=\sum_r \sum_s \left(\frac{p_r p_s |\mu_r-\mu_s|}{\sum_r \sum_s p_r p_s |\mu_r-\mu_s|} \right) I_{rs} \equiv \sum_r \sum_s w_{rs} I_{rs}; \tag{7}$$

where the weights w_{rs} are non-negative and sum to unity, with $w_{rr}=w_{ss}=0$ by construction. Thus $\hat{S}(1)$ may be interpreted as a weighted average identification index like $S(0)$ but with pairwise weights equal to shares in the total health gap between regions $NA(1)$ rather than in the population *N*. $\hat{S}(1)$ is a unit free measure that will take a maximum value of one if there is no overlapping of regional health distributions and richer regions are invariably healthier on average. More generally, $\hat{S}(1)$ is equal to $S(0)$ plus $\operatorname{cov}(I_{rs}, |\mu_s-\mu_r|/A(1))$, where the covariance may again be expected to be positive if richer regions tend to be healthier than poor ones.

$S(1)$ has the same units as *H* and is invariant to translations of the health outcome measure. With only two regions, the reduction in the IRHS gap caused by a unit increase in one person's health would again be greatest for inhabitants of the poorer region with health outcomes equal to the modal level in the richer region. And, more generally, it will also be the case that increasing the health of inhabitants of the poorest region may not necessarily have the most impact on the IRHS gap given that alienation is a linear function of the mean health gap. Assuming that the marginal cost of health improvements is independent of initial health then the minimum feasible cost of eliminating alienation through a targeted policy of health improvements would be proportional to the sum over all but the healthiest region of the product of region population and the mean health gap with the healthiest region, i.e. to $\sum_{r \neq R} n_r \left(\mu_R - \mu_r \right)$ if the richest region is also the healthiest.

3. Empirical analysis

The new class of measures is used to examine the evolution of income-related health differences between the regions of Great Britain. Our empirical analysis employs data on self-reported health from waves 1 to 18 of the British Household Panel Survey (BHPS; University of Essex, Institute for Social and Economic Research, 2010)), covering the years 1991 through 2008.

Established in 1991, the BHPS was a panel household survey with yearly interviews of all adults in each household covering a range of topics including health, work, education, income, family, and social life. The BHPS was 'primarily designed to be representative of all persons who are resident in Britain at multiple time points corresponding to the waves of data collection' (Lynn, 2006). In particular, the original sample was intended to be representative of all persons resident in private households in Great Britain (England, Scotland (south of the Caledonian Canal) and Wales) in 1991, with 13840 persons of all ages identified in wave 1 as original sample members at 8167 selected addresses (Lynn, 2006). The BHPS was permanently boosted at wave 9 in 1999 by the recruitment of new extension samples in Wales and (the whole of) Scotland, with a target sample size in each country of 1500 households in order to permit independent analysis of the two countries. The BHPS was replaced by the successor study, Understanding Society, following wave 18 in 2008.

The study is based on NUTS 1 statistical regions – Wales, Scotland and the Government Office Regions in England. The ordering of regions by mean income could be determined using income data from the BHPS, but instead we rely primarily on information about living standards from Households Below Average Income (HBAI) as this is generally considered to be the most reliable source of evidence on UK household net income and poverty (Office for National Statistics, 2015a).

Sample weights are used throughout the analysis with these being given by standardised BHPS cross-sectional respondent weights for each wave, where the standardisation takes account of wave-specific regional differences in population structure by sex and five year age band compared to Great Britain as a whole. Standard errors for all inequality and stratification measures are generated using a bootstrap procedure in which re-sampling is carried out at the cluster (Primary Sampling Unit) rather than individual level within each stratum, reflecting the sample design.[7]

3.1 Regional ordering by income

Estimation of income-related health stratification indices requires the prior ordering of regions by income. For all years from 1995 onwards, we use HBAI statistics (Department of Work and Pensions, 2015) on mean equivalised household incomes before housing costs by region as the primary ordering criterion and, in the few cases of ties, further rank regions on the basis of the distribution of individuals by (Great Britain) quintile groups as reported in Regional Trends (Office for National Statistics, various years). HBAI for this period uses data from the Family Resources Survey to construct estimates of the total weekly income of all household members after deductions of income tax and other contributions but before housing costs, with this total being equivalised to take account of the size and composition of households. Three-year centred moving averages are reported at the regional level as single-year estimates are considered too volatile. For the years prior to the introduction of the HBAI series, we use, as the closest comparable measure calculable from published statistics, three

[7] See Biewen (2002) on the use of bootstrap inference for inequality and mobility measurement.

year averages of normal weekly disposable per capita household income by (standard statistical) region based on Family Expenditure Survey data.

Table 1 reports the ordering of NUTS 1 regions by income over the study period. Within England, the North-South divide is clear, with the North-East consistently the poorest region in the country while London and the South East were the most prosperous throughout. Living standards in Wales were comparable to those of the poorest English regions, while those in Scotland, though higher, were still lower on average than in England as a whole in all years. Rankings between 1991 and 1994 appear to be broadly consistent with those from 1995 onwards, with the possible exceptions of those for the East and West Midlands.

Table 1: Ordering of NUTS 1 regions by income

Region Year	North East	North West	Yorks & Humber	East Midlands	West Midlands	East of England	London	South East	South West	Wales	Scotland
1991	1	6	3	7	4	8	11	10	9	2	5
1992	1	6	4	7	3	8	11	10	9	2	5
1993	2	5	4	7	3	8	11	10	9	1	6
1994	1	5	4	8	3	7	11	10	9	2	6
1995	1	4	3	5	7	9	10	11	8	2	6
1996	1	4	3	5	7	9	10	11	8	2	6
1997	1	4	3	5	7	9	10	11	8	2	6
1998	1	4	3	5	7	9	11	10	6	2	8
1999	1	4	3	5	6	9	11	10	7	2	8
2000	1	4	3	5	6	9	11	10	7	2	8
2001	1	4	6	3	5	9	11	10	8	2	7
2002	1	6	5	4	3	9	11	10	8	2	7
2003	1	6	3	5	4	9	11	10	8	2	7
2004	1	5	4	6	3	9	11	10	8	2	7
2005	1	6	3	5	2	9	11	10	8	4	7
2006	1	6	4	5	2	9	11	10	8	3	7
2007	1	6	4	3	2	9	11	10	8	5	7
2008	1	3	4	6	2	9	11	10	7	5	8

Source: own calculations, from HBAI, Regional Trends and Family Spending statistics. Regions are ranked in ascending order from the poorest (1) to the richest (11).

3.2 General health and health-related quality of life (HRQoL) variables

Respondents have been asked in all waves of the BHPS about their general state of health, but there are differences in the phrasing of this question between waves. Thus in all waves except wave 9, respondents were explicitly asked to think about their health over the past 12 months compared to people of their own age and say whether on the whole it had been very poor, poor, fair, good or excellent (BHPS variable: HLSTAT). In contrast, respondents were simply asked in wave 9 to say in general whether their health is poor, fair, good, very good or excellent, with this question also asked in wave 14 (BHPS variable: HLSF1). Ordinal measures of self-assessed health status have been widely used in the health economics literature to explore the relationship between health and income (see, e.g. O'Donnell et al., 2015). To make the interpretation of results more intuitive, we reverse the numerical coding of the BHPS variables so that higher scores correspond to better health.

The general health question asked in waves 9 and 14 is the first item in the Short Form (SF) health survey, with version 1 of the 36 item questionnaire administered in both waves (SF36: BHPS variables HLSF1-HLSF10D). The SF health survey is designed to measure functional health and well-being from the individual's point of view and is widely used in clinical trials (see Ware, 1993). Responses to SF-36 may be used to estimate measures of health-related quality of life (HRQoL) using a SF-6D preference-based algorithm (Brazier et al. 2002). The resultant cardinal HRQoL measure is bounded in the unit interval with full health corresponding to a value of one, the worst possible health outcome for anyone alive resulting in a score of 0.301, and death implicitly assigned a value of zero.

3.3 Empirical results

The results for the general health and HRQoL variables are discussed in turn, with the HRQoL analysis providing both headcount and gap indices but only for the two years in which the SF Health Survey was administered as part of the BHPS.

3.3.1 General health

Table 2 presents the main results from the analysis of headcount IRHS by NUTS 1 region for the two ordinal measures of general health, HLSTAT and HLSF1. Figure 1 plots the estimates of the normalised headcount index $\hat{S}(0)$, together with the associated 95% confidence intervals based on the bootstrap standard errors. The HLSTAT results suggest a slight upward trend in normalised headcount IRHS, rising from about 0.035 in the early 1990's to roughly 0.045 by 2008, while the two HLSF1 estimates for 1999 and 2004 both lie at the top end of this range. Thus both sets of estimates indicate that population health is better in richer regions since the index values are positive. More specifically, the normalised index gives the mean difference in the probabilities that the self-assessed health of a randomly chosen inhabitant of a richer region is better rather than worse

Figure 1. General health normalised headcount indices $\hat{S}(0)$ by NUTS 1 region, 1991-2008

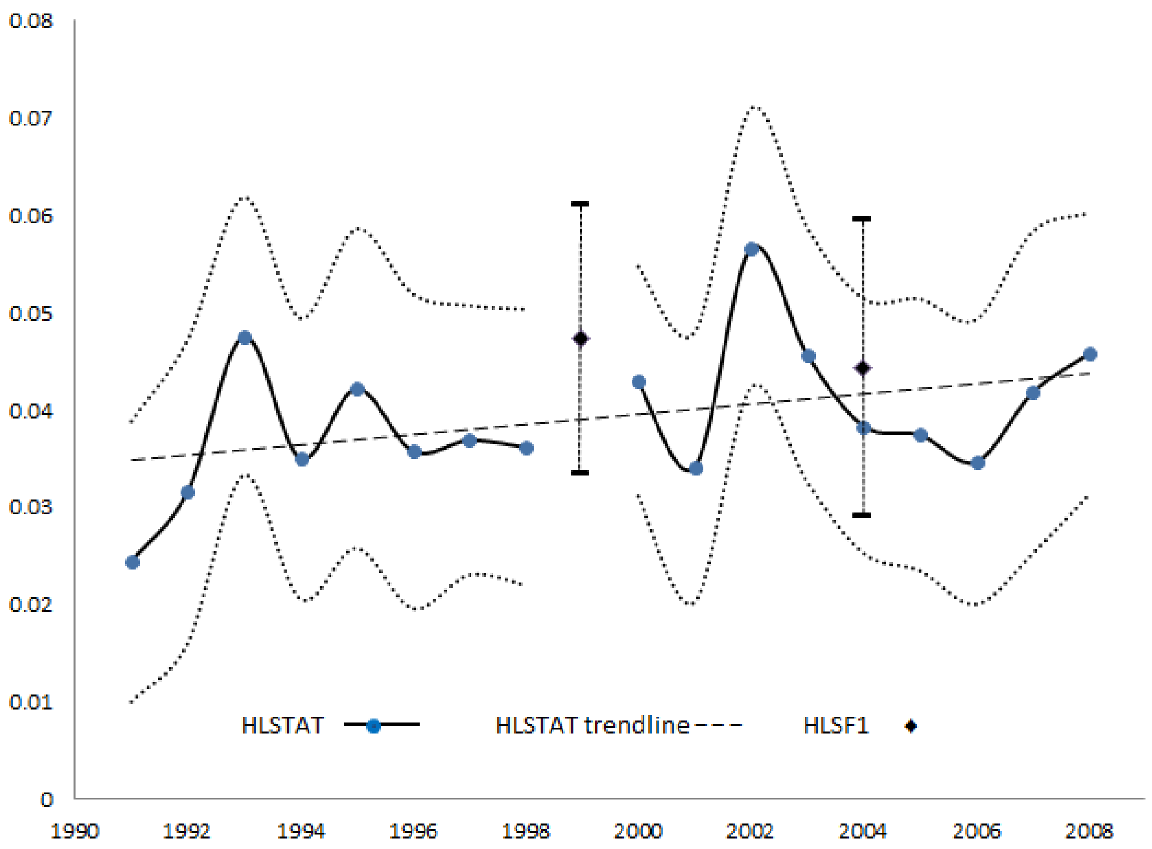


Table 2. General health headcount IRHS indices, 1991-2008

Year	Headcount index S(0)		Normalised headcount index $\widehat{S}(0)$		$\widehat{S}(0)$ based on 1991 population structure	
	HLSTAT	*HLSF1*	*HLSTAT*	*HLSF1*	*HLSTAT*	*HLSF1*
1991	0.0219 **		0.0244 **		0.0244 **	
	0.0066		*0.0074*		*0.0073*	
1992	0.0284 **		0.0315 **		0.0350 **	
	0.0072		*0.0080*		*0.0081*	
1993	0.0428 **		0.0475 **		0.0514 **	
	0.0066		*0.0073*		*0.0080*	
1994	0.0314 **		0.0349 **		0.0407 **	
	0.0066		*0.0074*		*0.0078*	
1995	0.0379 **		0.0421 **		0.0524 **	
	0.0075		*0.0084*		*0.0100*	
1996	0.0321 **		0.0357 **		0.0451 **	
	0.0074		*0.0083*		*0.0089*	
1997	0.0331 **		0.0368 **		0.0423 **	
	0.0064		*0.0071*		*0.0084*	
1998	0.0325 **		0.0361 **		0.0419 **	
	0.0065		*0.0072*		*0.0087*	
1999	-	0.0426 **	-	0.0473 **	-	0.0473 **
		0.0064		*0.0070*		*0.0070*
2000	0.0386 **		0.0429 **		0.0423 **	
	0.0054		*0.0060*		*0.0068*	
2001	0.0307 **		0.0340 **		0.0327 **	
	0.0064		*0.0071*		*0.0078*	
2002	0.0510 **		0.0566 **		0.0562 **	
	0.0066		*0.0073*		*0.0073*	
2003	0.0411 **		0.0456 **		0.0382 **	
	0.0060		*0.0067*		*0.0071*	
2004	0.0345 **	0.0399 **	0.0383 **	0.0443 **	0.0351 **	0.0409 **
	0.0060	*0.0070*	*0.0067*	*0.0078*	*0.0070*	*0.0081*
2005	0.0337 **		0.0374 **		0.0389 **	
	0.0064		*0.0071*		*0.0080*	
2006	0.0311 **		0.0346 **		0.0339 **	
	0.0067		*0.0075*		*0.0090*	
2007	0.0375 **		0.0417 **		0.0408 **	
	0.0076		*0.0084*		*0.0090*	
2008	0.0411 **		0.0457 **		0.0412 **	
	0.0066		*0.0074*		*0.0085*	

*Source: Own calculations based on equation (4). Bootstrapped standard errors in italics based on 500 replications. Statistical significance at 1% and 5% levels are denoted by ** and * respectively.*

Table 3: Detailed breakdown of the HLSTAT headcount IRHS index by NUTS 1 Region in 2004

Region		Popn share %	Pairwise identification indices: $I_{row,col} = \text{sgn}(row - col)\left(\text{P}\left(H_{row} > H_{col}\right) - \text{P}\left(H_{col} > H_{row}\right)\right)$											Regional index $S_{row}(0)$	$S(0)$ Share %
			NE	WA	WM	YH	NW	EM	SC	SW	EE	SE	GL		
North East	NE	4.2	0	0.104**	0.088*	0.081*	0.133**	0.102**	0.157**	0.142**	0.182**	0.174**	0.120**	0.0054**	15.6
				0.029	*0.035*	*0.040*	*0.029*	*0.034*	*0.029*	*0.030*	*0.035*	*0.030*	*0.034*	*0.0011*	
Wales	WA	5.2		0	-0.021	-0.023	0.027	-0.005	0.052**	0.038*	0.071**	0.063**	0.013	0.0015**	4.4
					0.023	*0.027*	*0.019*	*0.020*	*0.017*	*0.019*	*0.021*	*0.018*	*0.019*	*0.0006*	
West Midlands	WM	8.3			0	-0.003	0.050*	0.016	0.076**	0.061*	0.096**	0.089**	0.035	0.0040**	11.7
						0.032	*0.025*	*0.026*	*0.024*	*0.026*	*0.029*	*0.024*	*0.026*	*0.0013*	
Yorks & Humber	YH	9.0				0	0.051	0.019	0.076**	0.062*	0.097**	0.089**	0.037	0.0044**	12.8
							0.030	*0.030*	*0.029*	*0.029*	*0.032*	*0.030*	*0.030*	*0.0014*	
North West	NW	12.2					0	-0.033	0.025	0.012	0.044	0.036	-0.015	0.0030**	8.8
								0.023	*0.020*	*0.021*	*0.024*	*0.021*	*0.022*	*0.0009*	
East Midlands	EM	8.4						0	0.059**	0.045*	0.079**	0.071**	0.019	0.0028**	8.1
									0.021	*0.021*	*0.027*	*0.023*	*0.023*	*0.0005*	
Scotland	SC	8.9							0	-0.013	0.017	0.009	-0.040*	0.0026**	7.4
										0.020	*0.023*	*0.020*	*0.020*	*0.0005*	
South West	SW	9.3								0	0.031	0.023	-0.026	0.0025**	7.3
											0.024	*0.022*	*0.023*	*0.0005*	
East of England	EE	10.3									0	-0.008	-0.059*	0.0039**	11.3
												0.024	*0.026*	*0.0010*	
South East	SE	15.0										0	-0.051*	0.0051**	14.7
													0.021	*0.0016*	
London	GL	9.1											0	-0.0007	-2.1
														0.0014	
$S(0)$														0.0345**	
														0.0060	

*Source: Own calculations. The matrix is symmetric about the leading diagonal. Bootstrapped standard errors in italics based on 500 replications. Statistical significance at 1% and 5% levels are denoted by ** and * respectively.*

than that of a randomly selected inhabitant of a poorer region, conditional on the two individuals being from different regions and controlling for any demographic differences between the populations of the two regions. Thus $\widehat{S}(0)$ values about 0.04 imply that if two individuals were chosen at random from the standardised populations of two different regions then the probability that the healthier of the pair would be from the richer region was roughly 52% (since (0.52 – (1−0.52)=0.04). While this may not be much better than evens, the difference is statistically significant in all years.

Further insight into the source of this income-related health stratification can be gained from Table 3, which presents detailed results for HLSTAT in 2004 that may be taken as being typical of those obtained for both health measures and all years. Regions are ordered from the poorest to the richest with the values in the main body of the table being the pairwise identification indices. Thus the {NE, EE} value of 0.182 implies that if two individuals were randomly chosen from the standardised populations of the North East and Eastern England then there was a 59.1% chance that the person reporting better health would be from Eastern England. Indeed, health on this yardstick was substantially worse in the North East, the poorest region in Great Britain, than in all other regions − as indicated by the string of large and significantly positive values in the {NE} row. As a result, the contribution of the North East to the overall headcount IRHS index value of 0.0345 was 15.6% despite the population share of the region being only 4.2%. Conversely self-rerported health in London, the richest region in Great Britain, was significantly worse than in a number of other prosperous British regions, leading to the (insignificantly) negative net contribution of London to $S(0)$.

One possible cause of the observed trend in normalised headcount IRHS is changes in the demographic structure of the British population over the study period. To explore the possible effects of demographic change on the results, the indices were re-estimated with the cross-sectional weights for all years standardised on the basis of the population structure in Great Britain in 1991. These 'fixed population structure' estimates are reported in the final pair of columns in Table 2 and suggest that headcount IRHS would have been virtually constant over the entire study period if it had not been for changes in the composition of the British population by sex and age class.

3.32 Health-related quality of life

Table 4 presents results from the analysis of income-related HRQoL stratification by NUTS 1 region in 1999 in 2004. The main estimates of the headcount indices reported in Panel A are all positive but appreciably smaller than the corresponding estimates in Table 2 for the general health measure HLSF1, which is also obtained from the SF health survey but is not used in the computation of the HRQoL measure. More specifically, the normalised index values $\widehat{S}(0)$ imply that the difference in the chances of a representative inhabitant of a richer region having a higher rather than a lower level of HRQoL than a representative inhabitant of a poorer region was only of the order of $2.5\% - 3\%$.

One possible cause of the discrepancy between the HLSF1 and HRQoL estimates is that the former is a discrete variable that can only take five possible values whereas the latter may be considered as a continuous variable over the (truncated) unit interval.[8] To explore the possible effect of discretization on the results, the indices were re-estimated with the HRQoL data recoded into five classes, where the class boundaries were chosen in the two waves such that the proportion of the British population falling into each class was the same as for the HLSF1 variable. The discretized estimates reported in Panel B are only marginally different from the main estimates, which may be taken to imply that the considerable difference

[8] In practice the number of discrete values that the health utility index can take is finite but it does run into several hundred.

between the HRQoL and HLSF1 results is not due to the categorical nature of the latter variable but rather reflects substantive differences in the constructs underlying the two measures of health status. Additional senstivity tests (results not reported) show that the headcount index is generally robust to the grouping of the HRQoL data into quantiles so long as there are no fewer than about 5 health classes.

Table 4. Selected HRQoL IRHS indices, 1999 and 2004

	Headcount indices		Gap indices	
	S(0)	Normalised $\widehat{S}(0)$	S(1)	Normalised $\widehat{S}(1)$
A: Main results				
1999	0.0222 **	0.0247 **	0.00049 **	0.0523 **
	0.0055	*0.0061*	*0.00014*	*0.0116*
2004	0.0275 **	0.0305 **	0.00052 **	0.0547 **
	0.0056	*0.0062*	*0.00016*	*0.0102*
B: Discretized HRQoL data				
1999	0.0239 **	0.0265 **		
	0.0065	*0.0072*		
2004	0.0280 **	0.0310 **		
	0.0055	*0.0060*		
C: "Pure" health stratification indices				
1999	0.0306 **	0.0340 **	0.00056 **	0.0592 **
	0.0057	*0.0064*	*0.00017*	*0.0112*
2004	0.0354 **	0.0394 **	0.00059 **	0.0621 **
	0.0068	*0.0076*	*0.00019*	*0.0113*
D: Regional mean health data				
1999			0.00618 **	
			0.00139	
2004			0.00821 **	
			0.00134	

*Source: Own calculations. Bootstrapped standard errors in italics based on 500 replications. Statistical significance at 1% and 5% levels are denoted by ** and * respectively.*

Table 5: Detailed breakdown of the HRQoL headcount IRHS index by NUTS 1 Region in 2004

Region		Popn share %	Pairwise identification indices: $I_{row,col} = \text{sgn}(row-col)\left(\text{P}\left(H_{row} > H_{col}\right) - \text{P}\left(H_{col} > H_{row}\right)\right)$											Regional index $S_{row}(0)$	$S(0)$ Share %
			NE	WA	WM	YH	NW	EM	SC	SW	EE	SE	GL		
North East	NE	4.2	0	0.048 ** *0.018*	0.099 ** *0.029*	0.025 *0.029*	0.087 ** *0.022*	0.042 *0.026*	0.111 ** *0.026*	0.118 ** *0.027*	0.103 ** *0.035*	0.104 ** *0.024*	0.128 ** *0.035*	0.0036 ** *0.0009*	13.1
Wales	WA	5.2		0	0.042 *0.021*	-0.024 *0.025*	0.035 * *0.014*	-0.011 *0.021*	0.058 ** *0.016*	0.059 ** *0.021*	0.046 *0.025*	0.045 * *0.023*	0.071 ** *0.024*	0.0018 ** *0.0006*	6.6
West Midlands	WM	8.3			0	-0.074 * *0.029*	-0.010 *0.024*	-0.059 ** *0.022*	0.020 *0.020*	0.016 *0.028*	0.001 *0.027*	-0.001 *0.028*	0.029 *0.033*	-0.0001 *0.0012*	-0.2
Yorks & Humber	YH	9.0				0	0.063 * *0.030*	0.015 *0.023*	0.087 ** *0.024*	0.093 ** *0.030*	0.077 * *0.035*	0.079 ** *0.026*	0.102 ** *0.029*	0.0043 ** *0.0012*	15.7
North West	NW	12.2					0	-0.048 *0.026*	0.026 *0.019*	0.027 *0.021*	0.012 *0.030*	0.011 *0.026*	0.039 *0.027*	0.0021 * *0.0009*	7.8
East Midlands	EM	8.4						0	0.075 ** *0.019*	0.078 ** *0.024*	0.061 * *0.031*	0.062 * *0.025*	0.089 ** *0.030*	0.0025 ** *0.0005*	9.0
Scotland	SC	8.9							0	-0.004 *0.020*	-0.016 *0.024*	-0.020 *0.027*	0.011 *0.023*	0.0020 ** *0.0004*	7.3
South West	SW	9.3								0	-0.016 *0.027*	-0.019 *0.032*	0.014 *0.033*	0.0022 ** *0.0005*	8.1
East of England	EE	10.3									0	-0.001 *0.041*	0.030 *0.040*	0.0021 *0.0009*	7.6
South East	SE	15.0										0	0.031 *0.033*	0.0029 *0.0024*	10.7
London	GL	9.1											0	0.0039 * *0.0022*	14.2
$S(0)$														0.0275 ** *0.0056*	

*Source: Own calculations based on equation (4). The matrix is symmetric about the leading diagonal. Bootstrapped standard errors in italics based on 500 replications. Statistical significance at 1% and 5% levels are denoted by ** and * respectively.*

Table 6: Mean HRQoL differences between NUTS 1 Regions in 2004

		HRQoL		Pairwise mean HRQoL differences: $(\mu_{col} - \mu_{row})$											Regional index
Region		Mean	Rank	NE	WA	WM	YH	NW	EM	SC	SW	EE	SE	GL	$A_{row}(1)$
North East	NE	0.785 **	1	0	0.006	0.022 **	0.008	0.021 **	0.011	0.022 **	0.029 **	0.027 **	0.029 **	0.030 **	0.00087 *
		0.005			*0.005*	*0.006*	*0.007*	*0.005*	*0.006*	*0.005*	*0.006*	*0.006*	*0.006*	*0.008*	*0.00037*
Wales	WA	0.791 **	2		0	0.016 **	0.002	0.015 **	0.005	0.016 **	0.022 **	0.021 **	0.022 **	0.024 **	0.00078 **
		0.003				*0.005*	*0.006*	*0.004*	*0.005*	*0.004*	*0.004*	*0.005*	*0.005*	*0.007*	*0.00029*
West Midlands	WM	0.806 **	6			0	-0.013	-0.001	-0.011 *	0.000	0.007	0.005	0.007	0.008	0.00120 **
		0.004					*0.007*	*0.005*	*0.006*	*0.005*	*0.005*	*0.006*	*0.006*	*0.008*	*0.00022*
Yorks & Humber	YH	0.793 **	3				0	0.013 *	0.002	0.014 *	0.020 **	0.018 *	0.020 **	0.021 *	0.00100 **
		0.006						*0.006*	*0.007*	*0.006*	*0.007*	*0.007*	*0.007*	*0.009*	*0.00019*
North West	NW	0.806 **	5					0	-0.010 *	0.001	0.008	0.006	0.007	0.009	0.00085 **
		0.003							*0.005*	*0.004*	*0.005*	*0.005*	*0.005*	*0.007*	*0.00015*
East Midlands	EM	0.795 **	4						0	0.011 *	0.018 **	0.016 **	0.018 **	0.019 *	0.00058 **
		0.004								*0.005*	*0.005*	*0.006*	*0.006*	*0.008*	*0.00020*
Scotland	SC	0.807 **	7							0	0.006	0.005	0.006	0.008	0.00061 **
		0.003									*0.005*	*0.005*	*0.005*	*0.007*	*0.00016*
South West	SW	0.813 **	10								0	-0.002	0.000	0.001	0.00077 **
		0.003										*0.006*	*0.005*	*0.007*	*0.00022*
East of England	EE	0.812 **	8									0	0.002	0.003	0.00120 **
		0.005											*0.006*	*0.008*	*0.00027*
South East	SE	0.813 **	9										0	0.001	0.00075 **
		0.004												*0.008*	*0.00023*
London	GL	0.815 **	11											0	0.00083 *
		0.007													*0.00037*
$A(1)$															0.00944 **
															0.00125

*Source: Own calculations The matrix is symmetric about the leading diagonal. Regional index is a weighted sum of the absolute mean health gaps. Bootstrapped standard errors in italics based on 500 replications. Statistical significance at 1% and 5% levels are denoted by ** and * respectively.*

Table 7: Detailed breakdown of the HRQoL IRHS gap index by NUTS 1 Regions in 2004

Region		Popn share %	100*Pairwise IRHS gap indices: $100{*}\,S_{row,col}=I_{row,col}\left|\mu_{col}-\mu_{row}\right|$											Regional index $S_{row}(1)$	$S(1)$ Share %
			NE	WA	WM	YH	NW	EM	SC	SW	EE	SE	GL		
North East	NE	4.2	0	0.0300	0.2146	0.0209	0.1845	0.0455	0.2467*	0.3379**	0.2770	0.2980**	0.3829	0.000088*	17.0
				0.0371	*0.1258*	*0.0600*	*0.0949*	*0.0582*	*0.1242*	*0.1270*	*0.1620*	*0.1086*	*0.2087*	*0.000036*	
Wales	WA	5.2			0.0648	-0.0053	0.0513	-0.0052	0.0925	0.1316	0.0958	0.0994	0.1680	0.000038**	7.3
					0.0522	*0.0185*	*0.0324*	*0.0145*	*0.0506*	*0.0773*	*0.0796*	*0.0625*	*0.0873*	*0.000013*	
West Midlands	WM	8.3				-0.0987	-0.0007	-0.0640	0.0008	0.0113	0.0006	-0.0009	0.0239	0.000001	0.2
						0.0806	*0.0198*	*0.0513*	*0.0192*	*0.0286*	*0.0316*	*0.0337*	*0.0574*	*0.000015*	
Yorks & Humber	YH	9.0					0.0794	0.0037	0.1195	0.1884	0.1423	0.1586	0.2198	0.000080	15.5
							0.0780	*0.0198*	*0.0830*	*0.1189*	*0.1328*	*0.0864*	*0.1267*	*0.000045*	
North West	NW	12.2						-0.0495	0.0029	0.0207	0.0071	0.0085	0.0346	0.000025*	4.8
								0.0550	*0.0186*	*0.0262*	*0.0368*	*0.0252*	*0.0496*	*0.000012*	
East Midlands	EM	8.4							0.0850	0.1385	0.0987	0.1091	0.1695	0.000045*	8.7
									0.0565	*0.0788*	*0.0987*	*0.0717*	*0.1169*	*0.000018*	
Scotland	SC	8.9								-0.0025	-0.0077	-0.0125	0.0084	0.000028*	5.4
										0.0126	*0.0211*	*0.0189*	*0.0217*	*0.000013*	
South West	SW	9.3									-0.0028	-0.0002	0.0018	0.000049*	9.4
											0.0196	*0.0265*	*0.0408*	*0.000020*	
East of England	EE	10.3										-0.0001	0.0089	0.000040	7.7
												0.0494	*0.0535*	*0.000027*	
South East	SE	15.0											0.0044	0.000062	12.1
													0.0409	*0.000039*	
London	GL	9.1											0	0.000061	11.9
														0.000050	
$S(1)$														0.000517**	
														0.000155	

*Source: Own calculations based on equation (4). The matrix is symmetric about the leading diagonal. Bootstrapped standard errors in italics based on 500 replications. Statistical significance at 1% and 5% levels are denoted by ** and * respectively.*

Table 5 presents the pairwise headcount IRHS indices for 2004, with the regions ordered by income as in Table 3. The regional pattern of identification is broadly similar to that for HLSTAT with HRQoL generally better in more prosperous regions, such that all but two of the significant estimates are positive. Comparison with Table 6, which reports mean differences in HRQoL between regions in 2004, shows that those regional pairs with significant identification indices also tended to have significant differences in mean HRQoL. In particular, the HRQoL of a representative North Easterner was likely to have been lower than that of representative individuals from virtually every other British region so it comes as little surprise that mean HRQoL in the North East was significantly less than in all other regions except Wales. Conversely, the regional HRQoL distributions of the four most prosperous British regions, including London, were statistically indistinguishable from each other and there was also no significant differences in mean HRQoL levels. Ordering regions by HRQoL rather than income yields a "pure" HRQoL headcount index of 0.0354 for 2004, as reported in Panel C of Table 4, compared to the income-related value of 0.0275.[9]

Table 4 also reports the IRHS gap index $S(1)$, which takes into account not only the degree of identification but also the absolute differences in mean HRQoL between regions. $S(1)$ may loosely be interpreted as a measure of the perceived average difference in HRQoL between regions given the degree of regional identification. If all regional HRQoL distributions were fully identified (e.g. if all individuals had the average HRQoL of their own region) then $S(1)$ would be equal to twice the between-region generalised concentration index, and if the ordering of regions by income was also identical to that by HRQoL then $S(1)$ would equal the mean absolute HRQoL gap as given by the alienation index $A(1)$. Thus, taking 2004 as an example, the IRHS gap $S(1)$ of 0.00052 QALYs may be compared to the counterfactual $S(1)$ value of 0.00821 reported in Panel D of Table 4 and the mean absolute

[9] The Spearman's rank correlation coefficient between the two sets of ranks is 0.909.

HRQoL gap of 0.00944 reported in Table 6, with the differences reflecting the loss of absolute between-region inequality (cf. Milanovic and Yitzhaki, 2002) due to the overlapping of regional HRQoL distributions and the imperfect correlation of regional HRQoL and income outcomes respectively. In both 1999 and 2004, regional HRQoL disparities were largely masked by the low degree of identification due to the overlapping of regional HRQoL distributions.

Table 7 presents the pairwise IRHS gap indices for 2004, where the signs are determined by the signs of the pairwise identification indices reported in Table 5. Comparison with Table 5 shows that those regions with the highest and lowest average HRQoL made an even larger contribution to the total value of $S(1)$ than to $S(0)$, with these broadly corresponding to the most and least prosperous regions. In particular, the North East accounted for one sixth of the total IRHS gap due to the substantially below-average level of HRQol in the region. The normalised gap measure $\hat{S}(1)$ reported in Table 4 provides a weighted average identification index like $S(0)$ but with pairwise weights equal to shares in the total HRQoL gap between regions $NA(1)$. The higher values of $\hat{S}(1)$ compared to $S(0)$ confirm the positive correlation between pairwise mean HRQoL gaps and identification indices, i.e. region pairs that formed more clearly defined regional strata in their combined HRQoL distribution also tended to have larger differences in mean HRQoL.

4. Conclusion

This paper offers a new class of income-related health stratification indices that may be used to both characterise and quantify differences in health outcomes between the regions of a country, where the socioeconomic dimension is taken into account by ranking the regions in terms of economic prosperity rather than population health status. One major attraction of the proposed approach is that that the degree of variation in health outcomes within as well as between regions is taken into account, unlike conventional methods for the measurement of between-region health inequality such as the social gradient. Nevertheless the methodology does not also take account of income variation within regions, with this remaining a topic for future research given evidence that health outcomes in the poorer regions of Britain are not only worse on average but also across the entire income distribution. For example, Marmot at al. (2010, Figure 2.10) shows that if one compares neighbourhoods with the same level of income deprivation then disabilty-free life expectancy is lower in the North East than in London at all levels of neighbourhood income deprivation.

More specifically, the proposed indices depend in general both on the degree to which the populations of different regions occupy well-defined layers or strata in the national distribution of the health outcome and on the scale of between-region differences in those outcomes if these are quantifiable. These two factors play the same role as alienation and identification respectively in the measurement of polarisation (cf. Duclos et al., 2004) though it is important to recognise that stratification is not the same as polarisation due to the different characterisations of identification employed in the two sets of measures. It should be evident that the measurement framework may also be used to analyse differences between population groups classified on the basis of class, gender or race rather than region.

The other major attraction of the proposed class of measures is their ease of interpretation and practical utility. In particular, the IRHS headcount or incidence index provides a measure that is equal to the population-weighted mean difference in the probabilities that the health of a randomly chosen inhabitant of a richer region is better rather than worse than that of a randomly selected inhabitant of a poorer region. This identification measure is well-defined even if only ordinal or qualitative data are available, which is often the case with survey measures of self-reported health, subjective well-being and life satisfaction. The IRHS gap index, if defined, further allows for the degree of alienation between regions and may be seen as a generalisation of conventional measures of inequality between regions. Each index is a population-weighted average of pairwise indices so it is possible to estimate the contribution of individual regions to overall levels of IRHS with the further potential to identify the characteristics or factors that contribute to stratification.

The new class of measures is used to examine the evolution of income-related health differences between the regions of Great Britain between 1991 and 2008, where it should be noted that the results are sensitive to the chosen level of spatial aggregation. In particular, aggregation over districts with widely differing levels of average income relative to the national average will tend to result in lower levels of IRHS. For example, a country-level analysis of IRHS between England, Wales and Scotland (results not reported) yielded insignificant estimates of both headcount and gap indices in virtually all years. Conversely, an analysis at the local authority district level would reveal localised pockets of both income deprivation and health disadvantage within regions that are masked in the current study based on regional average incomes.

The empirical findings reveal three main points of interest. First there is a significant positive association between regional health and income outcomes, with a randomly chosen inhabitant of a richer region more likely to have had both better rather than worse health than a randomly chosen inhabitant of a poorer region and better health on average. In particular, the North East stands out as having been both the poorest and least healthy region in Great Britain throughout the study period: for example the region accounted in 2004 for as much as 13.1% of HRQoL headcount IRHS and 17.0% of the HRQoL IRHS gap despite a population share of only 4.2%. Health outcomes were also significantly worse in Wales, Yorkshire & Humberside and East Midlands than in many of the more prosperous regions of Southern England, broadly supporting the notion of a North-South divide within England (cf. Whitehead 2014).

Second, regional differences in general health can not simply be inferred from the regional pattern of variation in life expectancy. In particular, levels of general health in Scotland were indistinguishable from similarly prosperous regions in the rest of Britain despite Scotland having had the lowest life expectancy of any region in Britain over the entire study period (National Records of Scotland, 2015; Office for National Statistics, 2015b), mirroring similar findings in Taulbut (2013) for West Central Scotland. However this individual result should not be taken to imply that regional levels of prosperity were more strongly correlated with general health status than with life expectancy.[10]

Third, the lack of any apparent trend in headcount IRHS after controlling for changes in the demographic structure of the British population points to the persistence of the root causes of the observed differences in general health between regions. Whitehead (2014, p.5) observes that these causes are the same across the country, resulting from differences between socioeconomic groups not only in terms of poverty but also in the power and resources needed for health, in exposure to health damaging environments and in opportunities to enjoy positive health factors and protective conditions. Additionally, population health in certain areas, most notably in Northern England, Wales and Scotland, may have continued to have been affected by the legacy of heavy industry and its decline.

The empirical study could be extended through the use of health data from Understanding Society, the successor study to the BHPS, though differences in the general health questions between the two studies would limit comparability considerably.[11] The adoption of a longitudinal study design would further allow for the inclusion of death as a

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[10] For example, the correlation of regional average income with average HRQoL and life expectancy was 0.493 and 0.574 respectively for men in 2004.

[11] In particular, Understanding Society does not include the variable HLSTAT and only contains version 2 of the 12 item SF health survey, potentially limiting interest to HLSF1.

separate health outcome category, providing the basis for an analysis of income-related stratification in healthy life expectancy. It would also be of interest to examine inter-regional differences in other life outcomes, such as subjective well-being, life satisfaction or educational attainment. More generally, the measurement framework could be implemented with sub-regional (e.g. super output area) statistics used in place of individual data in the construction of the regional health distributions.

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